Abstract

This paper is concerned with the relationship between wages and unemployment. Using UK regions and individuals as the basis for our analysis, the following questions are investigated. First, is the wage equation a relationship between unemployment and wages or wage changes? Second, can we identify the relationship completely by looking at regional wages and regional unemployment or do regional wages depend on aggregate unemployment as well? Third, are wages influenced only by the current state of the labour market or do contracts lead to wages depending on labour market conditions in the last boom or upon entry into the job? Finally is the wage-unemployment relationship influenced by inflation, competition or the housing market?

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Wage Equations, Wage Curves and All That

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Introduction

The relationship between wages and unemployment is one of the key building blocks in any comprehensive model of the macroeconomy and it is particularly important in understanding the diverse patterns of unemployment around the world. The publication of Blanchflower and Oswald's Wage Curve (Blanchflower and Oswald, 1994) gave a new lease of life to the study of this relationship, reinforcing the continuing movement of this study from the domain of aggregate time series modelling, where it started with Phillips (1958), to the domain of micro-data analysis.

A number of questions concerning the relationship between wages and unemployment remain unresolved. For example, is the relationship one between the level of wages and unemployment in the long run or is it a relationship between the change in wages and unemployment as originally specified by Phillips? How does the standard downward sloping relationship between wages and unemployment relate to the upward sloping cross-section relationship which tends to result from the condition that (identical) mobile individuals must be indifferent to where they live in the long run (based on Harris and Todaro, 1970)? Our purpose in this study is to shed some light on these and other questions and generally to undertake some clarification and synthesis. Our particular spur is the recent limited availability of the British New Earnings Survey dataset which provides extremely accurate individual information on hourly pay (provided by employers from payroll records) on a large sample of employees (around 1 percent of the employed population) on an annual basis since 1976 with a substantial panel element. The accuracy of the data and the length of the time series are of great value, particularly for the identification of dynamic patterns.

In what follows we address the following questions. First, is the basic wage equation a relationship between wage levels and unemployment or is it fundamentally a relationship between wage changes and unemployment (*ie* a Phillips Curve)? Second, can we identify the relationship completely by looking at the partial correlation between regional wages and regional unemployment or do regional wages depend also on conditions in the aggregate labour market and hence on aggregate unemployment? Third, are wages influenced only by the current state of the labour market or does the existence of risk-sharing employment contracts lead to workers being shielded from current labour market fluctuations and to individual wages depending on labour market conditions at the start of the job? Fourth, is the responsiveness of wages to unemployment attenuated at low levels of inflation, because of the downward rigidity of nominal wages? Fifth, does increased competition lead to a rise in the responsiveness of wages to current unemployment and a decline in the extent to which workers are shielded from labour market conditions? Finally, is it possible to identify a *positive* cross-section relationship between wages and unemployment if the equation is correctly specified? None of these questions are new; indeed all have been addressed as we shall see. Our aim is to provide some further persuasive answers.

The remainder of the paper is set out as follows. In the next section, we attempt to clarify the theoretical background, particularly with regard to the variety of potential relationships between regional wages and regional unemployment rates. Then, in Section 3, we discuss the data, an empirical strategy and the detailed questions which we hope to address. In Section 4, we present all the results and finish with some general conclusions.

2. Theoretical Background

In order to interpret the cross-section and time series relationships between wages and unemployment it is helpful to provide a simple model of the operation of regional labour markets. In this, we ignore effects due to nominal inertia in order to focus on the real aspects of the problem. So suppose we have an economy wide production function of the form

$$Y_t^{r} = f_t \sum_j a_{jt} N_{jt}^{r} \qquad (r < 1) \tag{1}$$

where Y_t = aggregate output (price = unity) and N_{jt} = employment in region j. t refers to time, $\mathbf{N_t}$ captures aggregate productivity effects, $\mathbf{"_{jt}}$ capture regional productivity effects and can, without loss of generality, be normalised to $\sum_j \alpha_{jt} = 1$.

The labour demand function based on (1) may be written as

$$W_{jt} = \alpha_{jt}\, \phi_t \left(N_{jt} / Y_t\right)^{1/\sigma} \tag{2}$$

where F = (1-D)$^{-1}$, the elasticity of substitution, and W_j is the (real) wage in region i. This may be rewritten as:

$$W_{jt} = \alpha_{jt} \left((1 - u_{jt})\, s_{jt}\right)^{-1/\sigma} X_t \tag{3}$$

where $u_{jt} = (L_{jt} - N_{jt}) / L_{jt}$, the regional unemployment rate, L_{jt} being the regional labour force, and $s_{jt} = L_{jt}/L_t$, the share of the total labour force in region j, L_t being the aggregate labour force. $X_t = \phi_t \left(Y_t / L_t\right)^{1/\sigma}$ is an aggregate productivity effect.

Turning now to the wage equation, this is the relationship between regional wages and labour market conditions which is described variously in the literature as the wage setting equation, the wage curve, quasi-labour supply equation etc. (See Blanchard and Katz (1997) for a general discussion.) This we specify as

$$W_{jt} = \gamma_{jt}\, f(u_{jt})\, X_t\, \omega_t \qquad (f' < 0) \tag{4}$$

where $(_{jt}$ reflects autonomous factors generating region specific wage pressure and T_t captures aggregate wage pressure factors measured relative to the aggregate productivity term X. Again, without loss of generality, $(_{jt}$ can be normalised so that $\sum_j \gamma_{jt} = 1$. Finally we have a migration equation which relates net migration, $\dot{s}_{jt}$, to relative wages W_{jt}/W_t, unemployment and regional amenities, a_{jt}. Thus we have

$$\dot{s}_{jt} = h\left(W_{jt} / W_t,\ u_{jt},\ a_{jt}\right) \tag{5}$$

where $h_1 > 0,\ h_2 < 0,\ h_3 > 0$.

Much of this paper is devoted to investigating the wage equation (4), but in the light of the recent cross-section and time series literature, exemplified by Blanchflower and Oswald (1994), it is worth commenting on the implications of this model. First note that in the short-run (s_j fixed), wages and unemployment will be determined by "supply" (*ie* the wage equation (4)) and demand (3) within each region with the aggregate productivity effect, X, influencing wages but not unemployment[1]. In the long-run, however, we can impose the zero migration condition $\left(\dot{s}_{jt} = o\right)$ and wages and unemployment will be determined by the wage equation (4) and the zero migration condition ((5) with $\dot{s}_{jt} = o)$. The labour demand equation (3) will then determine regional labour input, s_{jt}. The former two long-run curves can then be thought of as a downward sloping wage curve

$$W_j = g_j f(u_j) X w \tag{4'}$$

and an upward sloping zero migration condition

$$0 = h\left(W_j / W, u_j, a_j\right) \tag{5'}$$

Both curves are identified, respectively by regional variation in wage pressure, $(_j$, and regional variation in amenities, a_j. The former would include factors such as unionisation rates and industry rents and the latter would consist of factors like the weather, property prices and the beauty of the countryside.

In practice, of course, we do not see the long run. However, even if we take a single year cross section, we should still be able to identify the two curves so long as we can control adequately for regional variation in wage pressure in the case of $(4')$ and both amenities and migration in the case of (5)[2]. In practice, the problem is always going to be the adequacy of the cross-section control variables. This is particularly true when it comes to estimating the wage setting equation (4, $4'$) bearing in mind that we also have to control accurately for region specific labour quality when operating in the cross-section context. This suggests that the best way of identifying the wage setting equation is to use the time series variation in a region based panel when the only serious remaining difficulties will be to control adequately for the time variation in region specific wage pressure and labour quality. Examples of this strategy may be found in Jackman and Savouri (1991), Blackaby and Manning (1992), Blanchflower and Oswald (1994), Blanchard and Katz (1997), Bell (1997), (1997a) and it will be pursued here.

3. The Data and Empirical Strategy

In order to base an analysis of wage setting on the regional fixed effects model, it is essential that we have a measure of wages which is accurate and does not have an in-built mechanical relationship to the cycle[3]. This is particularly important in the time series context when we are attempting to estimate dynamic relationships which are so easily corrupted by

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[1] u_{jt} satisfies $a_{jt}\left((1 - u_{jt}) s_{jt}\right)^{-1/s} = g_{jt} f\left(u_{jt}\right) w_t$ which is independent of X. The model has, of course, been arranged to satisfy this sensible "neutrality" property (see Blanchard and Katz, 1997).

[2] The statement is somewhat at variance with the discussion on p.93 of The Wage Curve (Blanchflower and Oswald, 1994) where it is essentially argued that (5) can never by identified in a single cross-section because there is so much random time series variation in amenities. We do not find this argument persuasive.

[3] For example, weekly earnings data are strongly related to cyclical fluctuations in hours worked – indeed some of this relationship will remain even if weekly earnings are normalised on hours worked because of the convex relationship between weekly earnings and weekly hours generated by overtime premia and the like. Annual earnings data tend to suffer from even greater problems.

measurement error. The basic data source used here is the UK New Earnings Survey (1976-1997) which is a 1 percent sample survey of employees. The sampling frame is based on all individuals whose National Insurance (NI) number ends in the digits 14. Since NI numbers are issued to every individual prior to starting work and are retained for life, there is a large panel element in the data.

Complete data on earnings are provided for each individual and cover a specific week in April in each year. These data are provided by employers who are legally required to comply. The data cover hourly and weekly earnings plus detailed information on hours, overtime, age, occupation, industry, region and whether or not the individual was in the same job as in the previous year. The measure of wages which we use throughout is the weekly pay of those whose pay was unaffected by absence excluding overtime pay divided by weekly hours excluding overtime hours. The idea is to obtain a measure of hourly pay which excludes the overtime element in order to eliminate that part of pay which is explicitly sensitive to the business cycle. It is worth noting that because the data come directly from firms' payroll records, they are extremely accurate.

The only other individual variable worth commenting on is occupation category. This variable we use to divide individuals into four skill groups which serve as human capital variables in the absence of direct information on qualifications or schooling (see Data Appendix for details). Other data are described in the Data Appendix. We next turn to the method which we use to construct a regional data panel.

3.1 The construction of regional data

Our aim here is to generate a region based model in which wages are adjusted for individual composition effects. We do this in two ways. The first uses the panel nature of the data and estimates a fixed effects equation for each region. So the wage equation for region j is

$$\mathrm{w}_{\mathrm{ijt}} = a_{\mathrm{i}} + a_{\mathrm{jt}} + \sum_{\mathrm{k}} \mathrm{x}_{\mathrm{kijt}}\, b_{\mathrm{kj}} + e_{\mathrm{ijt}} \qquad (6)$$

where w is the logwage, i = individual, j = region, t = time, "$_{\mathrm{i}}$ is an individual effect, "$_{\mathrm{jt}}$ is a region specific time dummy, the x variables include age[4], age^2, tenure, four skill groups, nine industry dummies. Men and women are treated separately. Equation (6) is then estimated using the fixed effects estimator and the region specific time effects $\hat{a}_{jt}$ are used as composition corrected wages in the regional panel model. For the second method we estimate a separate cross-section regression for each year, pooling individuals across all regions. We include the same regressors as in the previous equation as well as regional dummies. So in each year we estimate

$$\mathrm{w}_{\mathrm{ijt}} = a_{\mathrm{ot}} + a_{\mathrm{jt}} + \sum_{\mathrm{k}} x_{\mathrm{kijt}} b_{\mathrm{kt}} + h_{\mathrm{ijt}} \qquad (7)$$

and again use $\hat{a}_{\mathrm{jt}}$ as the composition corrected wage in the regional panel. What is the difference between these two methods of composition correction? In the first, which we shall refer to as "first stage panel", the composition parameters, b_{kj}, differ across regions but remain constant through time. Furthermore, they are not subject to biases because of any

[4] The age variable is only identified because new individuals are continuously entering the sample. Otherwise it is completely absorbed by the fixed effects and time dummies. Of course, its interpretation as an age effect is conditional on the assumption that there is no cohort effect.

correlation between the x variables and unobserved individual effects. In the second method, which we term "first stage cross-section", the β parameters are the same across regions but differ through time. This allows for more accurate composition correction over a period where industry and skill effects have become increasingly dispersed. However, the β parameters will reflect any correlation between the x variables and the individual effects.

The consequences of these differences for the composition corrected wages are hard to specify precisely, a priori. However, it is clear that the increases in dispersion of wages across skills and industries that have taken place over the sample period, allied to the systematic differences in skill and industry mix across regions, will lead to greater regional divergence over time in composition corrected wages if we have constant coefficients than if the industry skill coefficients can vary over time.

At the second stage we use the region/time cells as the unit of observation. The basic equation has the form

$$\hat{\alpha}_{jt} = \omega_j + \omega_t + \beta_1 \hat{\alpha}_{jt-1} - \beta_2 \ln u_{jt} + \sum_{j=2}^{J} \left(\gamma_j D_j\right) t + \sum_k z_{kjt} \beta_{k+2} + \eta_{jt} \qquad (8)$$

u_{jt} is the regional unemployment rate, D_j is a region dummy, z_{kjt} are region variables. The inclusion of region specific time trends represents an attempt to capture systematic trends in region specific wage pressure, and the equation is dynamic in order to investigate the process of wage adjustment.

This latter issue has been the subject of some controversy in recent years with some arguing that β_1 is close to zero (*eg* Blanchflower and Oswald, 1994) and others reporting it as close to unity (*eg* Blanchard and Katz, 1997, 1999). In what follows, we propose to investigate a number of variants of (8) which we discuss below, as well as reporting on estimates of a dynamic wage equation based directly on individual data, making use of its panel properties. This essentially involves estimating (6) and (8) in one stage, using an equation of the form

$$w_{ijt} = \alpha_i + \omega_j + \omega_t + \beta_1 w_{ijt-1} - \beta_2 \ln u_{jt} + \sum_{j=2}^{J} \left(\gamma_j D_j\right) t + \sum_k x_{kijt} \beta_k' + \sum_k z_{kjt} \beta_{k+2} + v_{ijt} \quad (9)$$

where we include individual effects **(α_i)**, region effects **(ω_j)**, time dummies **(ω_t)** and regional trends as well as lagged wages, unemployment and other individual or regional controls. Because this equation contains region level variables, standard errors must be corrected for clustering on region (*ie* lack of independence of errors within regions – see Moulton, 1986).

3.2 Topics to be investigated

3.2.1. Dynamics

The fundamental question here, discussed at length in Black and Fitzroy (1996), Bell (1997a) and Blanchard and Katz (1997; 1999), is whether or not the wage setting equation really has the form of a Phillips Curve with β_1 = 1. Note that, despite appearances to the contrary, the time dummies, $\mathbf{T_t}$, take care of aggregate price normalisation, and aggregate price surprises, so (8) can be thought of as a real wage equation[5].

The first point to note is that the time dimension used in estimating (8) is 21 years, so Hurwicz type biases which arise in the estimation of dynamic fixed effects models (see Nickell, 1981) are going to be very small. In fact the key issue when it comes to estimating the coefficient on the lagged dependent variable, $\$_1$, is how well we can capture the effects of region specific variations in both unobserved labour quality and autonomous wage pressure such as those arising from variations in unionisation, rent capture, the extent of product market competition and so on. If these factors are not adequately captured and vary systematically both over time **and** differentially across regions, then our estimate of $\$_1$ will typically be seriously overstated. Since it is hard to generate actual variables to capture all these effects, it is vital to include region specific trends in order to try and control adequately for these factors.

3.2.2 Aggregate effects

While the use of time dummies enables us to control for all aggregate effects in estimating our regional wage setting model, we have enough time periods to enable us to make an attempt to disentangle the contents of the estimated time dummies (see Galiani, 1999). Why should we wish to do so? First, regional wages may be influenced not only by the state of the regional labour market but also by the state of the aggregate labour market. This may arise because some wages are, or were, set nationally or are influenced by the national situation. It is important to see if we can identify this effect because if it were significant, we would have regional wages being influenced by both regional and national unemployment (u_{jt} and u_t). Then, on aggregating, the total unemployment effect on national wages would be the sum of these u_{jt} and u_t effects and since this is what determines the operation of the national economy it is a topic worth pursuing.

A second important issue arises from the possibility that regional wages may depend not only on last period's wage in the same region but also on last period's aggregate wage. This may arise because of mobility or because of wage comparability factors when, in the process of wage bargaining, the parties to the bargain compare their situation with the national wage level for similar workers during the previous period. In order to pursue these questions, we estimate a simple aggregate model, making use of the estimated time dummies $\left(\hat{\omega}_t\right)$ from equation (8). The factors captured by these time dummies include the two variables already discussed, namely aggregate unemployment, ln u_t, and last period's aggregate wage. However now we are at the aggregate level, we must explicitly deal with the price level, in order to focus on real wages. So last period's aggregate wage should be the real wage $\left(w_{t-1} - p_{t-1}\right)$ where p is the ln (GDP deflator). Furthermore, given the dynamic structure of (8), we must normalised the dependent variable by subtracting $\left(p_t - \hat{b}_1\, p_{t-1}\right)$[6]. Additional variables required are a variable to capture nominal inertia or price surprises, one to capture autonomous fluctuations in aggregate wage pressure and a variable which reflects

[5] We also have information on regional cost of living differentials, so we can control for any effects on regional wages arising from variations in regional to aggregate cost of living differences.

[6] Note that (8) can be rewritten in real terms as

$\hat{a}_{jt} - p_t = w_j + \overline{w}_t + b_1\left(\hat{a}_{jt-1} - p_{t-1}\right)$ ……… So the time dummy in (8), $\mathbf{T_t}$, is equal to $\overline{w}_t + p_t - b_1 p_{t-1}$. So $w_t - \left(p_t - b_1 p_{t-1}\right) = \overline{w}_t$ = the specifically aggregate factors influencing real wages.

trend labour productivity. For the first of these, we use $\Delta^2 p_t$, for the second an estimate of the NAIRU, u_t^* and for the third, we simply use a smoothed version of aggregate labour productivity, π_t. So the equation we estimate is

$$\hat{w}_t - \left(p_t - \hat{b}_1 p_{t-1}\right) = g_0 + g_1\left(w_{t-1} - p_{t-1}\right) - g_2 \ln u_t + g_3 \ln u_t^* - g_4 \Delta^2 p_t + g_5 \mathit{p}_t + h_t \quad (10)$$

3.2.3 Contract effects

When firms are risk neutral and workers are risk averse and face high mobility costs, risk sharing contractual arrangements tend to arise which reduce the response of wages to current labour market shocks. Instead, we would observe wages being responsive to labour market conditions at the time when the employee was initially appointed, that is to “initial” unemployment rates (see, for example, Baker *et al* 1994, 1994a). Alternatively with low mobility costs, if the labour market gets very tight after a worker joins the firm, the employer is forced to respond in order to prevent the worker leaving. This leads to wages depending on the minimum unemployment rate since the worker started her job (see Beaudry and Di Nardo, 1991).

In order to investigate these issues we use the dynamic wage equation based directly on individual data (Equation 9). We do this because initial and minimum unemployment rates vary across individuals within regions even though they are based on regional unemployment rates, so there is more information at the individual level. Our strategy then is simply to include variables reflecting the regional unemployment rate at the start of the individual’s job or the minimum unemployment rate in the individual’s region since starting the job. In doing this, we face two problems. First, the start of a job in our data does not necessarily mean starting with a new employer – it means starting a new type of work. This is an inescapable limitation of our data. Second, if initial or minimum unemployment rates prove to be important, this may be subject to an alternative interpretation to the contractual one suggested above (see Bertrand, 1998). When unemployment is low, alternative employment is easy to find and employees may use this opportunity to improve the quality of their job matches. Consequently those jobs which start in periods of low unemployment may represent better job matches and thus pay higher wages, *ceteris paribus*. Furthermore, if employees pass through a period of low unemployment and remain with their employer, this suggests that they must have a particularly good match relative to the average and this again will generate an unwanted negative correlation between wages and the minimum unemployment rate. Furthermore, these correlations will not be eliminated by the inclusion of individual fixed effects. To pursue this issue we shall investigate the extent to which the quality of new matches does appear to fluctuate with the unemployment rate.

3.2.4. Miscellaneous questions

a) Inflation and wage rigidity

Here we investigate whether or not wages are less responsive to unemployment when inflation is low. The idea, following Card and Hyslop (1997), is that if nominal wages exhibit some degree of downward rigidity, then this will impede the response of wages to unemployment when inflation is low because more falls in nominal wages are “required” than when inflation is high. We pursue this issue simply by interacting the aggregate inflation rate with unemployment to see if the (absolute) unemployment coefficient is higher when inflation is high.

b) Wage responsiveness to unemployment and competition

Following Layard *et al* (1991; Chapter 4, Section 5), Nickell and Kong (1992) and Bertrand (1998), we pursue the question of whether wages become more responsive to current unemployment when firms operate in a more competitive environment. First, we simply investigate whether the unemployment effect is bigger in the second half of the period relative to the first half. This may be expected because of the significant decline in both product and labour market regulation in Britain in the early and mid 1980s which can be expected to have increased product market competition. Second, we investigate whether the unemployment effect is bigger for workers in the manufacturing sector, on the grounds that manufacturing is subject to a higher level of product market competition. Third, we follow Bertrand (1998) in investigating the effect of import penetration at the industry level on the impact of unemployment on wages. The argument here is that import penetration reflects the degree of competition in an industry.

c) The impact of mortgage payments on wages

The work of both Blackaby and Manning (1992) and Cameron and Muellbauer (1999) suggests that conditions in regional housing markets have an impact on regional wages. Here we focus on the impact of changes in mortgage payments on wages, the idea being that when interest rates, and hence mortgage payments, go up, this may increase upward pressure on wages. We do not pursue the question of whether house prices themselves have a direct impact on regional wages, basically because we see long-standing regional house price differentials as reflecting differences in amenities across regions. Since amenities are a crucial part of the migration condition (see (5)) and not of the wage equation (see (4)), their inclusion in the wage equation is arguably a misspecification which would tend to disrupt identification. We pursue this in our next item.

d) Can we identify an upward sloping zero-migration condition?

In our discussion of the theoretical background in Section 2, we note that we should be able to identify a positive cross-section relationship between regional wages and regional unemployment so long as we control for regional amenities (see equation 5′). The main (dis) amenities variable we use is an index of regional house prices which varies substantially across regions.

4. The Results

Here we present a sequence of results covering, in turn, the topics set out in the previous section, starting with the issue of dynamics.

4.1 The basic wage model and wage dynamics

In Table 1, we present some results on the standard regional model based on equation (8). In the first two columns we report results based in the first stage panel method for correcting regional wages for composition effects. The short-run elasticity of wages with respect to unemployment is –0.034 and the long-run elasticity is –0.13. The model is very simple but if we include the following additional variables in Column 1, we obtain: $\ln u_{jt-1}$, coefficient 0.0064 (t=0.9); proportion of unemployed with duration exceeding 12 months, coefficient

0.0046 (t=0.2); region specific consumer prices (rel to aggregate prices)[7], coefficients on both t, t-1 dated variables are negligible (t < 0.1). The same applies to region specific producer prices[8]. So, in future, we simple omit these additional variables. In Column 2, we see that the equation for women closely resembles that for men. This is true for much of what follows and since women move in and out of the sample substantially more frequently than men, we shall in future focus on men, to avoid excessive repetition. In Columns 3 and 4 we repeat the exercise using the first stage cross-section method to control for wage composition effects. For men, the short-run elasticity for wages with respect to unemployment is –0.039 and the long-run elasticity is –0.090. While the short-run elasticity in column 3 is similar to that in column 1, the long-run elasticities are a bit different essentially because the lagged dependent variable coefficient in the latter case is rather lower. We shall pursue this further below.

In Columns 5 and 6, we investigate the consequences of instrumenting unemployment on the grounds that it may itself be influenced by current wage shocks. This seems unlikely given the high degree of persistence in labour demand and the notoriously sluggish response of unemployment to shocks of any kind. Nevertheless we should pursue the issue, so in column 5 we use powerful, although perhaps unpersuasive instruments, namely the first and second lags on ln u, finding the results essentially unchanged. In Column 6 we use the most persuasive instrument we can find, namely that obtained by applying fixed UK region-specific industry weights to time varying US industry unemployment rates. This variable should be independent of UK regional wage shocks and, as we can see, using it as an instrument leaves the point estimate of the unemployment effect almost unchanged, although its standard error is dramatically increased, as we might expect. We conclude from this that it is safe to treat unemployment as exogenous.

Turning now to the thorny issue of wage dynamics, we can see from Columns 1 to 4 that the lagged dependent variable (LDV) coefficient is around 6 or 7 standard errors away from unity, suggesting that the equation will determine the level of wages in the long run[9]. Turning to Table 2, we see from Columns 1 and 2 that it is the inclusion of region specific trends that is responsible for this. When they are omitted, the lagged dependent variable coefficient moves up to 0.97 or 0.82 depending on which of the two methods described in section 2 is used to correct wages for composition effects. Given the order ($^1/_T$) downward Hurwicz bias generated in dynamic fixed effects models (see Nickell, 1981) the former figure (0.97) is almost certainly, and the latter figure (0.82) may well be, insignificantly different from unity even when T is 21.

So the key issue is whether or not we are justified in including region specific trends. It is worth noting that this is exactly the same situation as that ruling in the United States. The results reported on similar state based wage models in Bell (1997a) and Blanchard and Katz (1997) reveal an LDV coefficient in the absence of region or state time trends of 0.92 (se = 0.017). Bell (1997a) then shows in Table 6 that including trends for the 9 Census region reduces the LDV coefficient to 0.83 (se = 0.04) and including trends for all 51 states reduces it further to 0.56 (se = 0.04). In the British case, when wages are composition corrected using the panel method, the regional trends indicate that over a ten year period, wages in the outer reaches of Britain fall relative to those in the South East by between 2 percent (East Midlands) and 5.6 percent (north), holding constant age, skill and industry. If

[7] These prices are based on data collected by Regional Rewards Survey Limited, who sample approximately 100 British towns. The data used exclude housing costs.
[8] These prices are based on industry output prices which are converted to a regional basis by using fixed weights generated by the average proportion of sales in each industry in the region concerned.
[9] Of course, under the null of a unit coefficient on the lagged dependent variable, the relevant "t ratio" will not be distributed as t but a 6 standard error gap should enable us to reject the null.

we correct for composition using the cross-section method, these regional trend effects are about half the size. This is because, over the relevant period, skill and industry wage dispersion have been increasing. When we composition correct using the cross-section method, this increasing wage dispersion is captured by the increasing dispersion of the skill and industry coefficients in successive cross-sections. So it does not feed through into the regional trends. When we use the panel method, skill and industry coefficients remain constant and the increasing dispersion in these (as well as other) dimensions passes through into the region specific trends. Either way, the resulting pattern of trends is consistent with region-specific shifts in union power (declined in the Northern regions relative to those in the South), shifts in financial sector rents (risen in the South East) and the migration of higher quality workers to the South East. If we look at the individual based fixed effects regression in column 3 (based on equation (9) and a sample where each individual has 17 or more consecutive wage observations), this tells the same kind of story as far was wage dynamics are concerned as the regressions in Table 1 but it should control for the migration of high quality workers. In fact, the regional trends have the same pattern and order of magnitude as those in the region based panel with the first stage, cross-section method of correcting for composition.

To summarise, therefore, the inclusion of region-specific trends has some theoretical foundation (see the discussion in Section 1) and the estimated trends are consistent with casual priors. However, recent simulations by Autor and Staiger (1999; Table 5) indicate that even if state log wage means are generated by random walks (*ie* true LDV coefficient = 1), the estimated LDV coefficient in a similar model (50 states, 20 years) including state dummies and trends would be around 0.65 and even lower in the presence of measurement error[10]. Pursuing this further, we split our individual panel into regions and then estimate dynamic fixed effect wage equations within each region including time dummies as well as skill and industry dummies. Again, each sample is restricted to men for whom we have 17 or more consecutive wage observations. The resulting LDV coefficients are between 0.43 and 0.50 with standard errors not exceeding 0.005. With these so far away from unity, it is hard to imagine that these individual based equations would generate region specific random walks when aggregated[11], except in the unlikely event that the individual wages depend very heavily on lagged aggregate regional wages which are, of course, hidden away in the time dummies. We can pursue this a little further by taking the estimated time dummies, appropriately subtracting out the aggregate price level and then seeing if these are correlated with lagged regional (real) wages. When we do this, we find no significant relationship, so we conclude finally that the regional random walk model is implausible on the basis of our evidence.

Another way of looking at this whole question is to consider the cointegration properties of the regressions in Table 1. This is worth doing since it is obvious that regional wages have a unit root. So we consider the unit root properties of the residuals of the first four regressions in Table 1 estimated with the lagged dependent variable removed. If we can reject the hypothesis that the residuals from these regressions have a unit root, this will add

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[10] Such a world has the rather unappealing property that the cross-region variance of log wages would tend to increase without bound. Of course, the same is true of a model with different region specific trends but the argument here is that these are included to capture unobserved sample specific regional movements in labour quality and wage pressure, with no presumption that these would continue indefinitely.

[11] When estimating these individual based dynamic models we restrict the sample to those who are present for at least 17 consecutive periods in order to estimate the dynamics accurately. Of course, the aggregate wage contains these individual plus all those who are omitted because, for one reason or another, they have breaks in their wage observations. In fact, it would be reasonable to suppose that the wages of these individuals exhibit less persistence than those who remain in the sample. So if they were added in, they would tend to reduce the level of persistence in the aggregate.

force to our view that the long-run relationship is one between unemployment and the level of wages rather than between unemployment and the change in wages. To pursue this, we generate fitted residuals from the first four regressions in Table 1 <u>omitting lagged wages</u> and then consider unit root tests on the resulting panel. The tests we use are first, those described in Levin and Lin (1993) and second, those suggested in Maddala and Wu (1996). The Levin and Lin test involves estimating a model of the form

$$\Delta\, \mathrm{res}_{it} = d_{io} + d_1 \mathrm{res}_{it} + \mathrm{u}_{it}, \mathrm{l} = 1 \ldots \mathrm{N},\ \mathrm{t} = 1 \ldots \mathrm{T}.$$

and testing the null hypothesis $d_1 = 0$ against the alternative $d_1 < 0$, where res_{it} are the fitted residuals. With N=10, T=25, Levin and Lin provide a 5 percent critical value of the t statistic of –5.42. The actual set of t ratios generated by the equations corresponding to the first four columns of Table 1 are, successively, -3.94, -5.00, -6.54, -6.35. These suggest that we cannot reject the null hypothesis of a unit root in the residuals if wages are composition corrected using the panel method but we can reject the null in the case where wages are composition corrected using the cross-section method.

The Maddala-Wu test is somewhat more general than that of Levin and Lin since it allows for parameter heterogeneity in the fitted residual autoregressions. The procedure is first to estimate i=1 to N regressions of the form

$$\Delta\, \mathrm{res}_{it} = d_{io} + d_{i1} \mathrm{res}_{it} + \mathrm{u}_{it},\ \mathrm{t} = 1 \ldots \mathrm{T}$$

and then on each of these regressions perform a Dickey-Fuller unit root test using MacKinnon critical values, each with a P value = P_i. Maddala and Wu (1996) then show, following Fisher (1932), that under the null of a unit root, $-2 \sum_{i-1}^{N} \log \mathrm{P}_i \sim \mathrm{X}^2_{2N}$. This relies on their being no cross-region correlation, but since the fitted residuals are based on regressions including region and, more importantly, time effects, any remaining cross-region correlations should be minimal. The relevant 5 percent critical value when 2N=20 is 31.41 and the values of $-2\, \Sigma \log \mathrm{P}_i$ for the residuals corresponding to the first four models in Table 1 are successively 33.31, 34.20, 85.49, 73.15. In this case the null is rejected in all four cases. Overall, therefore, we are persuaded that the inclusion of the regional trends to capture unobserved region variables is a legitimate exercise and that the long-run relationship is between unemployment and the level of wages. The old fashioned Phillips type relationship is not consistent with these data.

4.2 The aggregate wage equation

In Table 3, we report on estimates of an aggregate wage equation based on (10). That is, we take the time dummies from the regional wage equation reported in Table 1, column 1, deflate them appropriately and then see if they are related to aggregate unemployment, lagged aggregate wages and other factors we expect to observe in an aggregate wage equation. There is no evidence here that there is either an aggregate lagged wage effect or an aggregate unemployment effect in the regional model. Consequently, the lagged wage and unemployment effects estimated in our region based or individual based models correspond directly to the economy-wide effects when the variables are aggregated up.

4.3 Contract effects

In Table 4 we report individual based regressions allowing wages to be influenced by the current unemployment rate and by either the unemployment rate at the start of the individual's job or the minimum unemployment rate since the start of the job. The results are clear cut. There is no impact of the initial unemployment rate but a strong negative effect of the minimum unemployment rate. Comparison with the individual based regression in Table 2 reveals that including the minimum unemployment rate leaves the current unemployment rate effect unchanged despite being of the same order of magnitude. This strong negative impact of the minimum unemployment rate is consistent with the existence of risk sharing contracts which lead to the impact of past favourable labour market shocks having a persistent effect, as discussed in Beaudry and Di Nardo (1991). However, as we noted in (3.2.3), this minimum unemployment rate effect may reflect the fact that employees remaining in the same job through a period of low unemployment may have a particularly good match and hence a "high" wage. To shed light on this question we investigate whether workers who change jobs during periods of low unemployment achieve better matches than normal. If so, the stayers in these periods will tend to be in better matches than normal. To pursue this, we split from the sample in each period those who stay in the same job and those who change jobs without an intervening spell of unemployment. We then take the median percentage year on year change in earnings in the former group from that in the latter group to obtain a measure of the match improvement obtained in each period by the movers. This series we then correlate with unemployment and obtain a correlation coefficient of –0.40 (P= 0.14). So we have a negative relationship significant at the 15 percent level, which at least hints that movers obtain better matches than normal in good times (low unemployment). So the alternative selection interpretation of the significant minimum unemployment effect cannot be rejected. We return to this issue later.

4.4 Inflation and wage rigidity

In Table 5, Column 1, we present a region-based equation where we interact inflation with log unemployment to see if there is any evidence that downward nominal wage rigidity reduces the impact of unemployment on wages when inflation is low (so that more falls in nominal wages are "required"). In fact the evidence reveals quite the reverse. If anything, the unemployment effect is smaller when inflation is high, thus when inflation is 10 percent ( ) p = 0.1), the short run unemployment elasticity is 0.031 as opposed to 0.041 when inflation is zero. A possible reason for this is that when inflation is high, relative wages across firms and sectors tend to fluctuate more randomly with inflation driven annual pay rises being negotiated at different times in different firms. This tends to make it harder for agents to appreciate the real market picture. Standard signal extraction arguments would then lead to a reduction in the effects of labour market tightness.

4.5 Unemployment effects and competition

The issue here is whether wages are more responsive to unemployment in a more competitive environment. As a first stab at investigating this question, we can make use of the fact that, during the early 1980s, many changes were introduced in Britain to reduce rigidities in both labour and product markets (*eg* restrictive anti-union legislation, privatisation) so that there was a distinct increase in the intensity of competition in both markets. So, in Table 5, column 2, we simply interact the unemployment effect with a dummy, D_t, which takes the value of unity after 1987. The result is that the short-run unemployment elasticity increases from

0.028 to 0.037, with the change being marginally significant. In column 3, we check that the inflation effect discussed above is not simply due to the fact that inflation was lower on average after 1987 than it was before. The evidence is that this is not the case. The inflation effect remains and it is much the same size both before and after 1987.

Returning to the interpretation of the significant effect of minimum unemployment on wages, under the risk sharing contract interpretation, increasing competition in product and labour markets should lead to a reduction in the minimum unemployment effect and a corresponding rise in the current unemployment effect (see Bertrand, 1988). Increased competitive intensity raises the cost of shielding workers from current market forces. In Table 6, column 1 we see that there is a huge increase in the current unemployment effect after 1987 alongside a substantial reduction in the minimum unemployment effect. Arguably this sustains both the contract interpretation of the minimum unemployment effect and the increased competition interpretation of the post-1987 dummy.

Continuing along the same lines, in Table 6, column 2, we interact the unemployment terms with a dummy which takes the value one if the individual works in the manufacturing sector. The idea here is that, very crudely, the manufacturing sector is more competitive than the remainder of the economy. Here we find that the unemployment effect in the manufacturing sector is significantly bigger than outside it, but the minimum unemployment effect is the same in manufacturing as elsewhere. Pursuing this issue in more detail, we follow Bertrand (1988) in using industry import penetration as an indicator of product market competition. Here we restrict ourselves to workers in the manufacturing sector, and to allow for the possibility that wage shocks have a direct impact on import penetration, we use the Revenga (1992) instrument, namely the source weighted real exchange rate associated with the relevant industry. The results set out in Table 7 show that industry import penetration has no impact either on wages or on the elasticity of wages with respect to unemployment. This is in stark contrast to the results reported in Bertrand (1988) where a 10 percentage point increase in import penetration raises the wage elasticity of unemployment by around two-thirds. It is, however, worth noting that Bertrand omits the lagged dependent variable, although this is unlikely to explain the substantial difference.

4.6 The impact of mortgage payments on wages

Our purpose here is to investigate the impact of household mortgage payments on wages to ascertain whether a rise in mortgage costs leads to upward pressure on wage demands. The variable we use is the average household mortgage debt in the region multiplied by the Building Society mortgage rate normalised by average household income in the region. Because of the possibility that wage shocks may impact on both mortgage debt and income, we use as an instrument the same variable but with mortgage debt and income (but not the mortgage rate) lagged two years. The results presented in Table 8 indicate that mortgage costs do indeed have an impact on wages. In the long run, a 5 percentage point increase in the share of mortgage payments in household income leads to a rise in wages of close to 2 percent. To give some idea of the practical consequences of this, if the Building Society mortgage rate rises from 7 percent to 10 percent, a not untypical increase during the early 1990s, this would lead to a long-run increase in wages of just over 1½ percent[12]. This represents a not insignificant increase in inflationary pressure.

[12] This is very close to the result implied by the equation 1 in Table 1 of Cameron and Muellbauer (1999). They include the Building Society Mortgage rate, and a 3 percentage point rise generates a 1.6 percent increase in wages in the long run.

4.7 Any upward sloping zero migration condition?

One of the most significant variations in amenities across regions is captured by the variation in the price of houses. Houses in the South East are at least twice the price of houses in the North, which then acts as a significant Northern amenity. If we control for this amenity, can we observe a long-run positive cross-section relationship between real wages and unemployment, reflecting the Harris-Todaro type equation (5′)? If we regress average real hourly wages in the region on unemployment and house prices averaged over the sample period, we obtain,

$$\ln \text{real wage}_j = \text{Const} + \underset{(0.078)}{0.22} \ln u_j + \underset{(0.101)}{0.58} \ln \text{house prices}_j.$$

$$\bar{R}^2 = 81,\ N = 10$$

So the answer is yes. While the number of regions is small, the relationship is strong enough to generate significant effects on both unemployment and house prices. Indeed, to show that this result emerges even in a single cross-section, if we undertake single cross-section regressions for 23 successive years and then average the outcomes, we obtain

$$\ln \text{real wage}_j = \text{Const.} + \underset{(0.085)}{0.17} \ln u_j + \underset{(0.107)}{0.49} \ln \text{house prices}_j$$

Then standard errors correspond to the average coefficient divided by the average t ratio.

5. Conclusions

1. The long-run elasticity of average regional wages with respect to regional unemployment is in the range 0.09 to 0.13 (Table 1, columns 1 and 3). The long-run elasticity of individual wages with respect to regional unemployment is around –0.053 (Table 2, Column 2).

2. While wages exhibit a high degree of autocorrelation both at the regional and the individual level, the lagged dependent variable coefficient is well below unity (between 0.5 and0.75). This relies on the inclusion of region specific trends but the balance of the evidence suggests that this is the correct specification, although they would not be required if all the relevant variables were available.

3. There is no evidence that regional wages, conditional on regional unemployment and lagged regional wages, depend on either aggregate unemployment or aggregate lagged wages. This suggests that the regional equation can be aggregated in a straightforward fashion and reflects the aggregate relationship[13].

4. There is strong evidence that individual wages are influenced both by the current unemployment rate and by the minimum unemployment rate since the individual's current job started. This suggests that risk sharing employment contracts may be important.

5. There is no evidence that the unemployment effect on wages is higher when inflation is high – indeed quite the reverse.

[13] Although note that because log wages depend on log unemployment, in the aggregate the log geometric mean of regional wages will depend on log aggregate unemployment plus a term in the cross-region dispersion of unemployment rates (see Layard *et al*. 1991, Chapter 6).

6. The impact of current unemployment on individual wages has risen in the last decade. By contrast the impact of the minimum unemployment rate since the current job started has diminished in the same period. This is interpreted as the effect of an increase in product and labour market competition. However, were have been unable to detect any interaction between the unemployment effect and import penetration.

7. There is a significant impact of average mortgage payments on regional wages. A rise in Building Society mortgage rates of 3 percentage points leads to a long-run increase in wages of just over 1½ percent.

8. It is possible to identify a positive cross-section relationship between wages and unemployment controlling for regional house prices.

Table 1: Regional Wage Equations, 1976-97
Dependent Variable: ln $wage_{jt}$

	1	2	3	4	5	6
	First stage, panel		**First stage, cross section**		**First stage, panel**	
	Men	**Women**	**Men**	**Women**	**Men**	**Men**
					IV	IV
ln $wage_{jt-1}$	0.730	0.679	0.568	0.570	0.760	0.619
	(0.044)	(0.056)	(0.061)	(0.068)	(0.046)	(0.11)
ln u_{jt}	-0.0342	-0.030	-0.039	-0.035	-0.0344	-0.0373
	(0.0041)	(0.0056)	(0.0064)	(0.0084)	(0.0051	(0.022)
Adjusted wage	T	T	T	T	T	T
Time dummies	T	T	T	T	T	T
Region dummies	T	T	T	T	T	T
Region trends	T	T	T	T	T	T
N	10	10	10	10	10	10
NT	210	210	210	210	210	190
Standard error	0.0049	0.0060	0.0076	0.0091	0.0050	0.0047

Notes:

i) Standard errors in parentheses. These are robust against heteroskedasticity.

ii) Equations 1, 2, 3 & 4 are estimated by OLS. In equations 5 and 6, ln u is treated as endogenous. In equation 5, ln u_{jt-1}, ln u_{jt-2} are used as instruments. In equation 6 the instrument is constructed by taking US industry specific unemployment rates which are used to create a British region specific index by using fixed, employment based, region specific industry weights. These data are only available up to 1995, at present.

iii) Wages are adjusted for composition effects by the mechanisms described in Section 3.1.

Table 2: Wages Dynamics, 1977-97, Men
Dependent Variable: ln wage

	1	2	3
	Region-based		Individual based
	First stage, panel	First stage, cross section	
ln $wage_{-1}$	0.970	0.820	0.540
	(0.0096)	(0.037)	(0.0087)
ln u_{jt}	-0.0239	-0.0233	-0.0245
	(0.0037)	(0.0056)	(0.0070)
Adjusted wage	T	T	.
Time dummies	T	T	T
Region dummies	T	T	T
Region trends	V	V	T
Individual dummies			T
N	10	10	6352
NT	210	210	116312
Standard error	0.0054	0.0081	

Notes:

i) Standard errors in parentheses. These are robust against heteroskedasticity and, in equation 2, they are adjusted for clustering on region (i.e lack of independence within region).

ii) In equation 1 & 2, regional wages are adjusted for composition effects using the two methods described in Section 2. In equation 3, all the relevant composition variables (*ie* skill, industry dummies) are included – since they are time varying, they are all identified in the fixed effects context.

iii) In equation 3, the sample is selected to include only men who appear in the sample for at least 17 consecutive years, so as to minimise the small T bias on the lagged dependent variable coefficient.

Table 3: Aggregate Wage Equation, 1977-97, Men (Equation 10)
Dependent Variable: $\hat{\omega}_t - (p_t - 0.73p_{t-1})$
(Based on Table 1, Column1)

	1	**2**
Const.	-3.31	-3.80
$(w_{t-1} - p_{t-1})$	-0.055	-0.10
	(0.19)	(0.20)
$\ln u_t$	0.024	0.013
	(0.019)	(0.023)
$\ln u_t^*$	0.0032	0.018
	(0.051)	(0.054)
$\Delta^2 p_t$	-0.33	-0.64
	(0.82)	(0.91)
π_t	0.22	0.26
	(0.11)	(0.11)
$\Delta \ln u_t$		0.018
		(0.021)
Standard error	0.014	0.014
$\bar{R}^2$	0.78	0.78
T	20	20

Notes:

i) These equations are based on (9). $\hat{\omega}_t$ is the estimated time dummy for period t in Table 1, Column 1, and 0.73 is the LDV coefficient in the same equation. w is the cross region average of the adjusted wages used as the dependent variable in Table 1, Column 1. p is the ln (GDP deflator). u_t is aggregate unemployment rate. u^* is the equilibrium rate. π is trend productivity.

ii) Standard errors in parentheses.

Table 4: Initial and Minimum Unemployment Effects, 1977-97, Men
Dependent Variable: ln $wage_{it}$
Individual Based Regressions

	1	**2**
ln $wage_{it-1}$	0.509	0.508
	(0.010)	(0.010)
ln u_{jt}	-0.0273	-0.0245
	(0.0067)	(0.0067)
ln initial u_{jt}	-0.00092	
	(0.0024)	
ln min u_{jt}		-0.0249
		(0.0026)
Skill, industry dummies	T	T
Time dummies	T	T
Region dummies	T	T
Individual dummies	T	T
Region trends	T	T
N	7841	7841
NT	128817	128817
Standard error		

Notes:
i) Standard errors in parentheses are robust against heteroskedasticity and are adjusted for clustering on region.
ii) Sample selected to include only men who appear in the sample for at least 15 consecutive years.

Table 5: Regional Wage Equations with Inflation Interactions, 1977-97, Men
Dependent Variable: ln $wage_{jt}$

	1	2	3
ln $wages_{jt-1}$	0.769	0.744	0.744
	(0.044)	(0.045)	(0.047)
ln u_{jt}	-0.0406	-0.0280	-0.0349
	(0.0053)	(0.0051)	(0.0058)
Δp_t x ln u_{jt}	0.0923		0.115
	(0.047)		(0.049)
D_t x ln u_{jt}		-0.0087	-0.0096
		(0.0044)	(0.0096)
D_t x Δp_t x ln u_{jt}			-0.0169
			(0.13)
Adjusted wage	T	T	T
Time dummies	T	T	T
Region dummies	T	T	T
Region trends	T	T	T
N	10	10	10
NT	210	210	210
Standard error	0.0049	0.0049	0.0049

Notes:
i) Robust standard errors in parentheses
ii))p is the inflation rate. D_t is a dummy which takes the value 1 from 1988 onwards, zero otherwise.
iii) Wages adjusted for composition effects using the first stage panel method described in Section 3.1. If we use the first stage cross-section method, the pattern of results is much the same.

Table 6: Minimum Unemployment Effects and Time and Manufacturing Interactions, 1977-97, Men
Dependent Variable: ln $wage_{it}$
Individual Based Regression

	1	**2**
ln $wage_{it-1}$	0.508	0.538
	(0.010)	(0.0086)
ln u_{jt}	-0.0024	-0.0169
	(0.0063)	(0.0080)
D_t x ln u_{jt}	-0.0252	
	(0.0062)	
M_{it} x ln u_{jt}		-0.0243
		(0.0041)
ln min. u_{jt}	-0.0298	-0.0191
	(0.0032)	(0.0040)
D_t x ln min. u_{jt}	0.0136	
	(0.0058)	
M_{it} x ln min. u_{jt}		0.00087
		(0.0028)
Skill, industry dummies	T	T
Time dummies	T	T
Region dummies	T	T
Individual dummies	T	T
Region trends	T	T
N	7841	6352
NT	128817	116312
Standard error		

Notes:
i) Standard errors in parentheses are robust against heteroskedasticity and are adjusted for clustering on regions.
ii) Sample selected to include only men who appear in the sample for at least 15 consecutive years.

Table 7: Import Penetration Effects, 1977-97, Men in Manufacturing
Dependent Variable: ln wage$_{it}$
Individual Based Regression: Manufacturing Only

	1	**2**
		IV
ln wage$_{it-1}$	0.473	0.473
	(0.013)	(0.0046)
ln u_{jt}	-0.0212	-0.0243
	(0.0093)	(0.0071)
import penetration $_{kt}$	0.0090	-0.041
	(0.055)	(0.18)
import penetration $_{kt}$ x lnu_{jt}	-0.0082	0.0091
	(0.021)	(0.073)
Skill, industry dummies	T	T
Time dummies	T	T
Region dummies	T	T
Individual dummies	T	T
Region Trends	T	T
N	2120	2120
NT	34622	34622
Standard error		

Notes:
i. i refers to individual, j to his region, k to his industry.
ii. Standard errors in parentheses are robust against heteroskedasticity and are adjusted for clustering on regions.
iii. Sample selected to include only men who appear in the sample for at least 15 consecutive years.
iv. The instruments are the source country weighted real exchange rate corresponding to the appropriate industry, alone and interacted with ln u.

Table 8: Mortgage Payment Effects, 1977-95, Men
Dependent Variable: ln $wage_{jt}$

	IV
ln $wage_{jt-1}$	0.519
	(0.065)
ln u_{jt}	-0.0319
	(0.0045)
average mortgage $cost_{jt}$	0.192
	(0.062)
Adjusted wage	T
Time dummies	T
Region dummies	T
Region Trends	T
N	10
NT	190
Standard error	0.0047

Notes:
i. Robust standard errors in parentheses
ii. Average mortgage $cost_{jt}$ = (average mortgage $debt_{jt}$ x building society mortgage $rate_t$) ÷ (average $income)_{jt}$. The instrument = (average mortgage debt $_{jt-2}$ x Building Society mortgage $rate_t$) ÷ (average income) $_{jt-2}$
iii. Wages adjusted for composition effects using the first stage panel method described in Section 3.1. If we use the first stage cross-section method, the results are very similar.

Data Appendix
Skill levels based on the Standard Occupation Classification

Skill Level	Major Groups	Constituent Minor Groups (2 digit)
Level 4	Managers and administrators (excluding office managers and managers/proprietors in agriculture and services). Professional occupations	10, 11, 12, 15, 19 20-27, 29
Level 3	Office Managers and managers/proprietors in agriculture and services Associate professional and technical occupations Craft and relations occupations Buyers, brokers, sales reps	13, 14, 16, 17 30-39 50-59
Level 2	Clerical, secretarial occupations Personal and protective service occupations Sales occupations (except buyers, brokers, sales reps) Plant and machine operatives Other occupations in agriculture forestry, fishing	40-46, 49 60-67, 69 72,73, 79 80-89 90
Level 1	Other elementary occupations	91-95, 99

Source: Elias (1995)

Other Variables

Wages: hourly wages of individuals in full-time employment whose earnings were not affected by absence. Computed by taking weekly earnings excluding overtime earnings divided by weekly hours excluding overtime hours. Reported by employers on a given week in April each year. New Earnings Survey (UK Office of National Statistics).

Unemployment**:** Unadjusted regional employment rates (April figure); UK Department of Employment Gazette (now called Labour Market Trends) and UK Regional Trends.

US unemployment: unemployment rate by industry of last job; Employment and Earnings (Bureau of Labour Statistics). GB Regional figures are obtained by weighting each industry by the proportion of employment of the industry in the GB region. UK Business Monitor.

Regional consumer prices: obtained using the percentage comparison of regional consumer prices to national consumer prices, and the national Retail Price Index; The Reward Group.

Regional producer prices: obtained using producer price index numbers of output (home sales) by industry. Annual Abstract of Statistics. Regional figures are obtained by weighting each industry by the proportion of employment of the industry in the region. Business Monitor.

National unemployment; claimant count: UK Labour Market Trends

Retail price index: UK Economic Trends

Trend productivity: smoothed (log GDP – log employment). Smoothed by fitting to a 5th order polynomial in time. UK Economic Trends.

Initial unemployment: Unemployment rate when the current job started (a new job starts when the individual declares not to be in the same job as 12 months before, and the unemployment rate of year preceding that survey date is used). Note that jobs at first entry into the survey are excluded because we cannot recover the starting date of the job and therefore the unemployment rate at the time.

Minimum unemployment: Lowest unemployment rate since the current job started. As above, note that jobs at first entry into the survey are excluded because we cannot recover the starting date of the job and therefore the unemployment rate over the missing years of tenure.

Import penetration: import penetration for products of the manufacturing industry. UK National Accounts (Blue Book)

Source based weighted real exchange rate: weighted average of the log real exchange rates of importing countries. The weights are the share of each country's imports in total imports in each industry in a base period (1985). Real exchange rates are obtained using nominal exchange rates and UK and foreign country RPI. Source: nominal exchange rates from the OECD Bilateral Trade Dataset, RPI are taken from International Financial Statistics of the International Monetary Fund.

Average mortgage cost: obtained by multiplying the average mortgage debt by Building Societies average Mortgage Rate. The average mortgage debt (avmrtg) is obtained as follows:

avmrtg(i)=[mort(i)/(avfte(i)*pop(i))]

mort(i)=stock of mortgage debt in region i.
avfte(i)=average earnings of men working full time (weekly)
pop(i)=population

The stock of mortgage debt is obtained as follows. First we estimate the size of the average mortgage by region (am),

am(i)=(xrmp/100)*(mortuk/fnluk)

xrmp is average mortgage payment in region i relative to the UK.
mortuk is mortgage stock in the UK
fnluk is the number of mortgages in the UK

Thus the size of the average mortgage by region is obtained scaling up or down the size of the average mortgage in the UK using the ratio of mortgage payments in a region relative to the UK.

The size of the average mortgage per region (am) is then multiplied by the number of mortgages per region (fnl) to obtain the stock of mortgage debt per region.

mort(i)=am(i)*fnl(i).

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